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AUTHOR	Lord, Frederic M.
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Frederic M. Lord

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Standard Error of an Equating by Item Response Theory

Abstract

A formula is derived for the asymptotic standard error of a true-score equating by item response theory. The equating method is applicable when the two tests to be equated are administered to different groups along with an 'anchor test.' Numerical standard errors are shown for an actual equating 1) comparing the standard errors of IRT, linear, and equipercentile methods; 2) illustrating the effect of the length of the anchor test on the standard error of the equating.

Standard Error of an Equating by Item Response Theory*

In item response theory (IRT), an examinee's expected number-right score ξ on test X is equal to the test characteristic function evaluated at the examinee's ability level θ :

$$\xi = \sum_{g=1}^{n_x} P_g(\theta) \tag{1'}$$

where $P_i(\theta)$ is the item response function, the probability of a correct answer to item i at ability level θ . If we have a second test, Y , measuring the same ability as X , the expected number-right score η on this test may be written as

$$\eta = \sum_{h=1}^{n_y} P_h(\theta) \quad . \tag{4'}$$

Equations (1') and (4') are parametric equations for the functional relationship between ξ and η . Note that this relationship is an exact mathematical one, not a statistical association. Given any θ , (1') and (4') determine a pair of values, ξ and η , that represent the same ability level as θ . Pairs of values (ξ,η) determined in this way are equated. In practice, it is often assumed that the functional relationship of η to ξ given by (1') and (4') can also be applied to actual number-right scores on the two tests, producing an equating of these scores.

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Here, we simply deal with the sampling errors in estimating the equating relationship of η to ξ . In (1') and (4'), estimated item parameters must be used. These are the source of the sampling errors in IRT equating. Note that the ability estimates for individual examinees are not used in (1') and (4') and thus will not appear in our formulas. Until now, the sampling errors of IRT equatings have never been estimated.

Data

In IRT equating, we frequently have a set of common items that are administered to all examinees. These are needed in order to get Test Y item parameters on the same scale as Test X item parameters. If the common items are external to tests X and Y , as assumed here, the common items are called the anchor test, or, in the present report, Test W . The sampling variance formulas to be obtained here can be modified in obvious ways for the case where some or all of the common items are internal to the tests that are being equated.

Designate the examinees who took both Tests X and W as Group 1; designate the examinees who took Tests Y and W as Group 2. Typically, every examinee falls in one of these two groups.

In practice when there is a series of test forms A,B,...,X,Y,Z,... (say), the 'Group 1' data on Test X are processed as soon as they become available in order to equate Test X to the preceding form. When the Group 2 data become available at some later date, it is often considered uneconomical to rerun the Group 1 data, so Group 2 is

run by itself. This case, where item parameters for Groups 1 and 2 are estimated separately, is the case to be considered here. (The simplifying assumption that is used below to approximate the sampling variances of the estimated item parameters is not available in the alternative case where Groups 1 and 2 are pooled and all parameters estimated simultaneously.)

New Equating Formulas

When parameters are estimated separately for groups 1 and 2, the item parameters and θ in (4') have a different origin and scale from the item parameters and θ in (1'). It is thus no longer possible simply to eliminate θ from (1') and (4') to obtain the relation of η to ξ. The customary procedure in this situation is to use the anchor test to transform the Group 2 item parameters on to the scale of the Group 1 item parameters. This procedure adds to the sampling variance of the transformed item parameters and greatly complicates any determination of the sampling variance of the subsequent equating. The procedures and formulas given below avoid this problem since they avoid any transformation of item parameters.

Equations (1') and (4') remain unchanged except that additional subscripts (explained below) are used. In particular, the symbols θ_1 and θ_2 must be distinguished because groups 1 and 2 use different ability scales:

$$\xi = \sum_g P_{g1}(\theta_1) \quad , \tag{1}$$

$$\eta = \sum_g P_{g4}(\theta_2) \quad . \tag{4}$$

The item response functions here are written P_{gp} where $p = 1,2,3,4$ refers to (test X, group 1), (test W, group 1), (test W, group 2), and (test Y, group 2) respectively, and $g = 1,2,\ldots,n_p$ where n_p, is the number of items in the appropriate test.

Let us write down similar equations for the expected number-right score ω on anchor test W:

$$\omega = \sum_g P_{g2}(\theta_1) \quad , \tag{2}$$

$$\omega = \sum_g P_{g3}(\theta_2) \quad . \tag{3}$$

The equation numbering keeps the tests in convenient order. The desired equation relation between η and ξ can be obtained by eliminating θ_1, θ_2, and ω from these four equations.

Computer programs are available for equating η to ξ by eliminating θ from (1') and (4'). These same programs can be used to equate ω to ξ in one step, using (1) and (2), then to equate η to ω in a second step using (3) and (4). This produces an equating of η to ξ for the presently relevant situation where Group 1 and Group 2 parameters are not on the same scale.

An estimated equating is obtained from (1) - (4) after replacing the true item parameters by their maximum likelihood estimates. Using carets to denote this change, we have

$$\hat{\xi} = \sum_g \hat{P}_{g1}(\theta_1) \quad , \tag{1"}$$

$$\hat{\omega} = \sum_g \hat{P}_{g2}(\theta_1) \quad , \tag{2"}$$

$$\hat{\omega} = \sum_g \hat{P}_{g3}(\theta_2) \quad , \tag{3"}$$

$$\hat{\eta} = \sum_g \hat{P}_{g4}(\theta_2) \quad . \tag{4"}$$

These equations show that $\hat{\eta}$ is a function of all the estimated item parameters together with the specified value of ξ .

Derivatives

For item g , instead of using a_g , b_g , and c_g to denote the three parameters commonly used in IRT, let us use t_{1gp} , t_{2gp} , and t_{3gp} , respectively. We will need certain derivatives for $r = 1,2,3$, obtained from (1")-(4"):

$$\frac{\partial \eta}{\partial t_{rg4}} = P_{g4}^{(r)}(\theta_2) \quad , \tag{5}$$

$$\frac{\partial \omega}{\partial t_{rg3}} = P_{g3}^{(r)}(\theta_2) \quad ,$$

$$\frac{\partial \omega}{\partial t_{rg2}} = P_{g2}^{(r)}(\theta_1) \quad ,$$

where $P_{gp}^{(r)}$ denotes the derivative of P_{gp} with respect to t_{rgp} .

Similarly,

$$\frac{\partial \eta}{\partial \theta_2} = \sum_g P'_{g4}(\theta_2) \quad ,$$

$$\frac{\partial \omega}{\partial \theta_1} = \sum_g P'_{g2}(\theta_1)$$

where P' denotes a derivative with respect to θ. Using the formula for the derivative of an implicit function, we also find from (1")-(4") for $r = 1,2,3$

$$\frac{\partial \theta_2}{\partial t_{rg3}} = - \frac{P^{(r)}_{g3}(\theta_2)}{\sum_g P'_{g3}(\theta_2)} \quad ,$$

$$\frac{\partial \theta_1}{\partial t_{rg1}} = - \frac{P^{(r)}_{g1}(\theta_1)}{\sum_g P'_{g1}(\theta_1)} \quad ,$$

$$\frac{\partial \theta_2}{\partial \omega} = \frac{1}{\sum_g P'_{g3}(\theta_2)} \quad .$$

Using the chain rule for derivatives, we find from the above formulas:

$$\frac{\partial \eta}{\partial t_{rg3}} = \frac{\partial \eta}{\partial \theta_2} \frac{\partial \theta_2}{\partial t_{rg3}} = -P^{(r)}_{g3}(\theta_2) \frac{\sum_g P'_{g4}(\theta_2)}{\sum_g P'_{g3}(\theta_2)} \quad , \qquad (6)$$

$$\frac{\partial \eta}{\partial t_{rg2}} = \frac{\partial \eta}{\partial \theta_2}\frac{\partial \theta_2}{\partial \omega}\frac{\partial \omega}{\partial t_{rg2}} = P_{g2}^{(r)}(\theta_1)\frac{\sum_g P'_{g4}(\theta_2)}{\sum_g P'_{g3}(\theta_2)} , \tag{7}$$

$$\frac{\partial \eta}{\partial t_{rg1}} = \frac{\partial \eta}{\partial \theta_2}\frac{\partial \theta_2}{\partial \omega}\frac{\partial \omega}{\partial \theta_1}\frac{\partial \theta_1}{\partial t_{rg1}} = -P_{g1}^{(r)}(\theta_1)\frac{\sum_g P'_{g2}(\theta_1)}{\sum_g P'_{g1}(\theta_1)}\frac{\sum_g P'_{g4}(\theta_2)}{\sum_g P'_{g3}(\theta_2)} . \tag{8}$$

Given ξ, we are now in a position to express $\hat{\eta}$ as a series in powers of $\hat{t}_{rgp} - t_{rgp}$ ($r = 1,2,3$; $g = 1,2,\ldots,n_p$; $p = 1,2,3,4$). We will write η'_{rgp} instead of $\partial \eta/\partial t_{rgp}$ and η''_{rgpshq} instead of $\partial^2\eta/\partial t_{rgp}\partial t_{shq}$.

$$\hat{\eta} = \eta + \sum_p \sum_g \sum_r (\hat{t}_{rgp} - t_{rgp})\eta'_{rgp}$$

$$+ \frac{1}{2}\sum_p \sum_q \sum_g \sum_h \sum_r \sum_s (\hat{t}_{rgp} - t_{rgp})(\hat{t}_{shq} - t_{shq})\eta''_{rgpshq} + \ldots \tag{9}$$

Sampling Variance

Transposing, squaring, and taking expectations, we find from (9) for fixed ξ,

$$\text{Var}\,\hat{\eta} = \mathcal{E}(\hat{\eta} - \eta)^2 = \sum_p \sum_q \sum_g \sum_h \sum_r \sum_s \eta'_{rgp}\eta'_{shq}\,\text{Cov}(\hat{t}_{rgp},\hat{t}_{shq}) + \ldots$$

When item parameters and abilities are both estimated simultaneously by maximum likelihood, it is not practical to use the usual sampling covariance formulas for all estimators simultaneously. As a rough approximation, it is customary (Lord, 1980, Section 12.3) to use instead the (simpler) formulas for the case where the ability parameters are known. We will use this rough approximation here to find $\mathrm{Cov}(\hat{t}_{rgp}, \hat{t}_{shq})$. Because of this approximation, our sampling variance of equating will be an underestimate.

In this case, all covariances involving two different items are exactly zero, as are all covariances involving a single item administered to two different groups of examinees. All nonzero variances and covariances are inversely proportional to N, the number of examinees. We now have

$$\mathrm{Var}\,\hat{\eta} = \sum_p \sum_g \Big[\sum_{r=1}^{3} \sum_{s=1}^{3} \{\eta'_{rgp}\eta'_{sgp}\,\mathrm{Cov}(\hat{t}_{rgp}, \hat{t}_{sgp})\}$$

$$+ \sum^{3}\sum^{3}\sum^{3} \{\cdot\} + \sum^{3}\sum^{3}\sum^{3}\sum^{3} \{\cdot\} + \ldots\Big] .$$

Some higher order terms are indicated here in order to make clear that the number of terms under summation signs does not increase too rapidly. The triple summation represents 3 times as many terms as the double summation, but each term in the triple summation is divided by $N^{3/2}$ whereas each term in the double summation is only divided by N. When N is several thousand, it is reasonable to expect that the higher order terms can be neglected, as is customary with asymptotic variances.

Our final asymptotic formula, then is

$$\mathrm{Var}\ \hat{\eta} \doteq \sum_{p=1}^{4} \sum_{g=1}^{n_p} \sum_{r=1}^{3} \sum_{s=1}^{3} \eta'_{rgp}\eta'_{sgp}\ \mathrm{Cov}(\hat{t}_{rgp},\hat{t}_{sgp}) \qquad (10)$$

The η' values required here are computed from (5) - (8). The covaraince are obtained by the usual formulas for covariances of maximum likelihood estimators of item parameters when ability parameters are fixed (Lord, 1980, p. 191).

Practical Application

Without data, it is difficult to make inferences about the magnitude of the sampling errors in IRT equating. Will they be larger or smaller than the sampling errors in conventional linear equating? In conventional equipercentile equating? Do sampling errors become large or small at extreme score levels?

Equation (10) has been applied to an equating of the Verbal score on the 90-item Form VSA4 of the Scholastic Aptitude Test (12/73 administration) to the 85-item Form XSA2 Verbal score (4/75 administration). All examinees took an SAT and also a 40-item anchor test. Petersen, Cook, and Stocking (1980) made separate LOGIST runs on the 130 items in the 1973 administration for a sample of 2665 examinees, and on the 125 items in the 1975 administration for a sample of 2686 examinees. They have allowed the use here of their item parameter estimates.

SAT scaled scores are a linear transformation of formula scores (rights minus one-quarter wrongs). Our results here are for the hypothetical case where all examinees answer all items. In this special case formula scores are a linear transformation of number-right scores, so scaled scores are likewise. Since a known linear transformation $A\xi + B$ of number-right scores ξ simply multiplies the standard error of $\hat{\eta}$ by the constant A, it is not difficult to obtain scaled-score standard errors from (10). A computer program to do this was written and run by Marilyn Wingersky.

For each of certain specified formula scores on XSA2, Table 1 shows 1) the equivalent scaled score found by the conventional linear procedure usually used for the SAT (Design IV A, Angoff, 1971), 2) the standard error of these equated (scaled) scores as found by the computer program AUTEST (Lord, 1975) assuming the validity of the linear model; also 3) the equivalent scaled score found by the IRT method of this report, and 4) the corresponding scaled-score standard error calculated from (10). The standard errors in Table 1 are best understood in comparison with the standard deviation of scaled scores, which is 106 for XSA2; and in comparison with the classical test theory standard error of measurement (due to imperfect test reliability), which is 31. Clearly the standard error of equating is small compared to the standard error of measurement.

Judging by the IRT standard errors, the equating is definitely nonlinear, at least outside the score range from 350 to 650. The IRT standard errors show a continued sharp increase as the minimum

Table 1

A Comparison of Linear and IRT Equatings and of Their Standard Errors

Selected formula scores*, XSA2	Linear Model		IRT Model	
	Equivalent scaled score	Standard error	Equivalent scaled score	Standard error
84	780	4.6	813.8	2.3
79.74	750	4.2	778.0	4.5
72.70	700	3.6	717.6	4.4
65.65	650	3.1	658.8	3.6
58.61	600	2.5	602.4	2.8
51.57	550	2.1	548.0	2.2
44.52	500	1.7	495.4	2.0
37.48	450	1.5	445.7	2.1
30.43	400	1.6	399.3	2.3
23.39	350	1.8	355.6	2.8
16.35	300	2.3	313.3	3.6
9.30	250	2.8	270.2	4.7
2.26	200	3.3	223.0	7.0
-5	150	3.9	163.5	15.6

*Although formula score is actually a discrete variable, it is for convenience treated here as continuous.

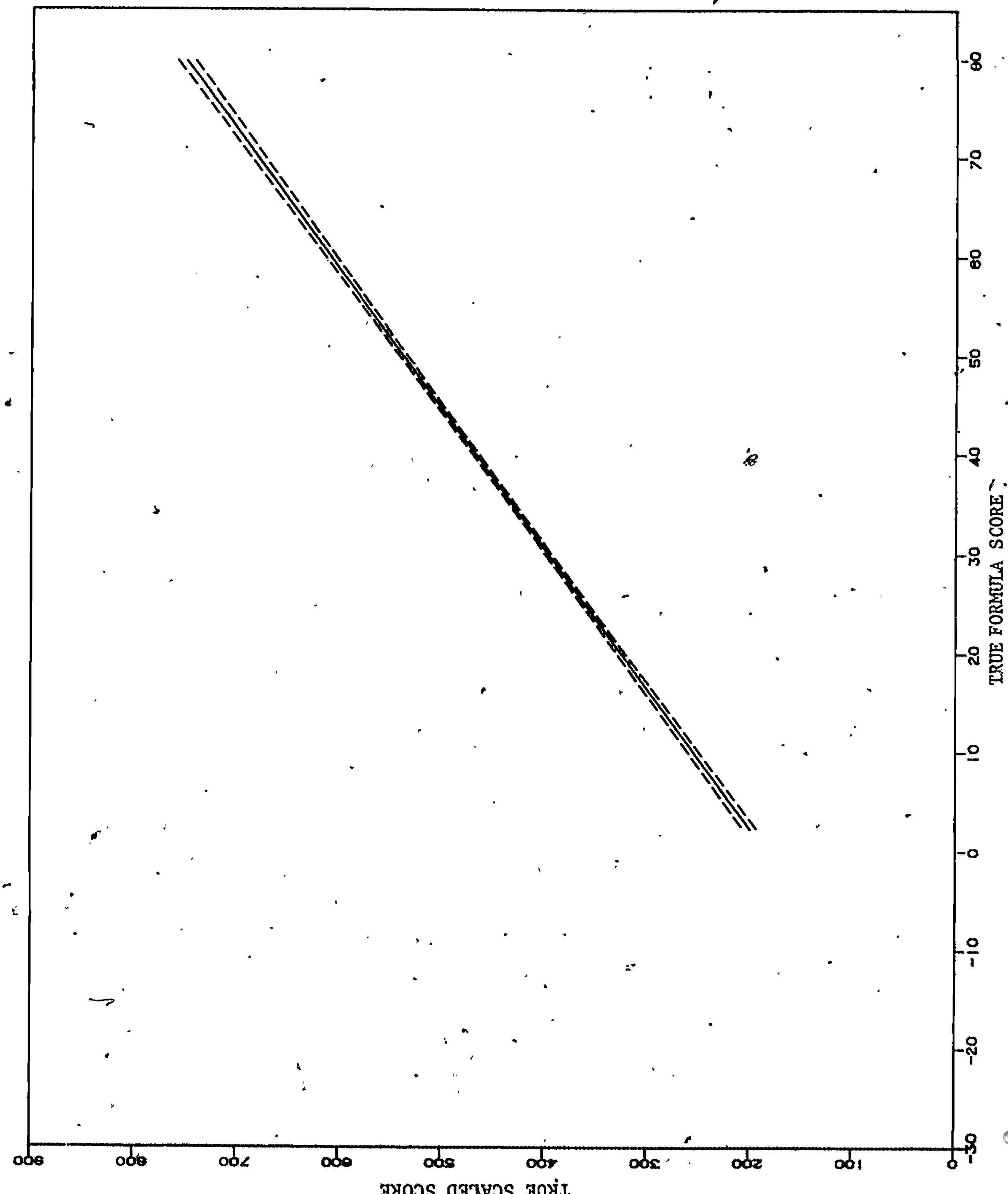
possible true formula score of -5.5 is approached. At the other end of the score scale, the IRT standard error increases up to a scaled score of 760 and decreases thereafter. The reason for the decrease at the upper end is that for a perfect score, the standard error of this kind of IRT equating is zero. Except at the upper end, the IRT standard error is larger than the linear.

The results of Table 1 are displayed in Figures 1-2. The straight line in Figure 1 shows the linear equating of true formula score on XSA2 to true scaled score on VSA4. The dashed lines are drawn two standard errors above and below the straight line.

Figure 2 similarly displays the curvilinear IRT equating of XSA2 to VSA4 and its standard error. The straight-line extension of the lower end of the equating (middle) line in Figure 2 was obtained by the method described in Lord (1980, pp. 210-211). It is shown in the figure for completeness, but no standard error is shown since there is no good theoretical basis for such an extension.

Table 2 compares present IRT equating with a conventional equipercentile equating of XSA2 to VSA4 via the anchor test. In conventional equating, an XSA2 score and a VSA4 score each equipercentile-ly equivalent to a given anchor test score are taken to be equivalent to each other. The standard error of the resulting equipercentile equating of XSA2 to VSA4 is given by $\sqrt{(SE^2_{XSA2} + SE^2_{VSA4})}$ where the SE under the radical sign are standard errors of separate equipercentile equatings of each test to the anchor test. Formulas for SE_{XSA2} and SE_{VSA4} are given in Lord (1981).

Figure 1. Linear equating of true formula score on XSA2 to true scaled score on VSA4. Dashed lines are two scaled-score standard errors above and below equating line.



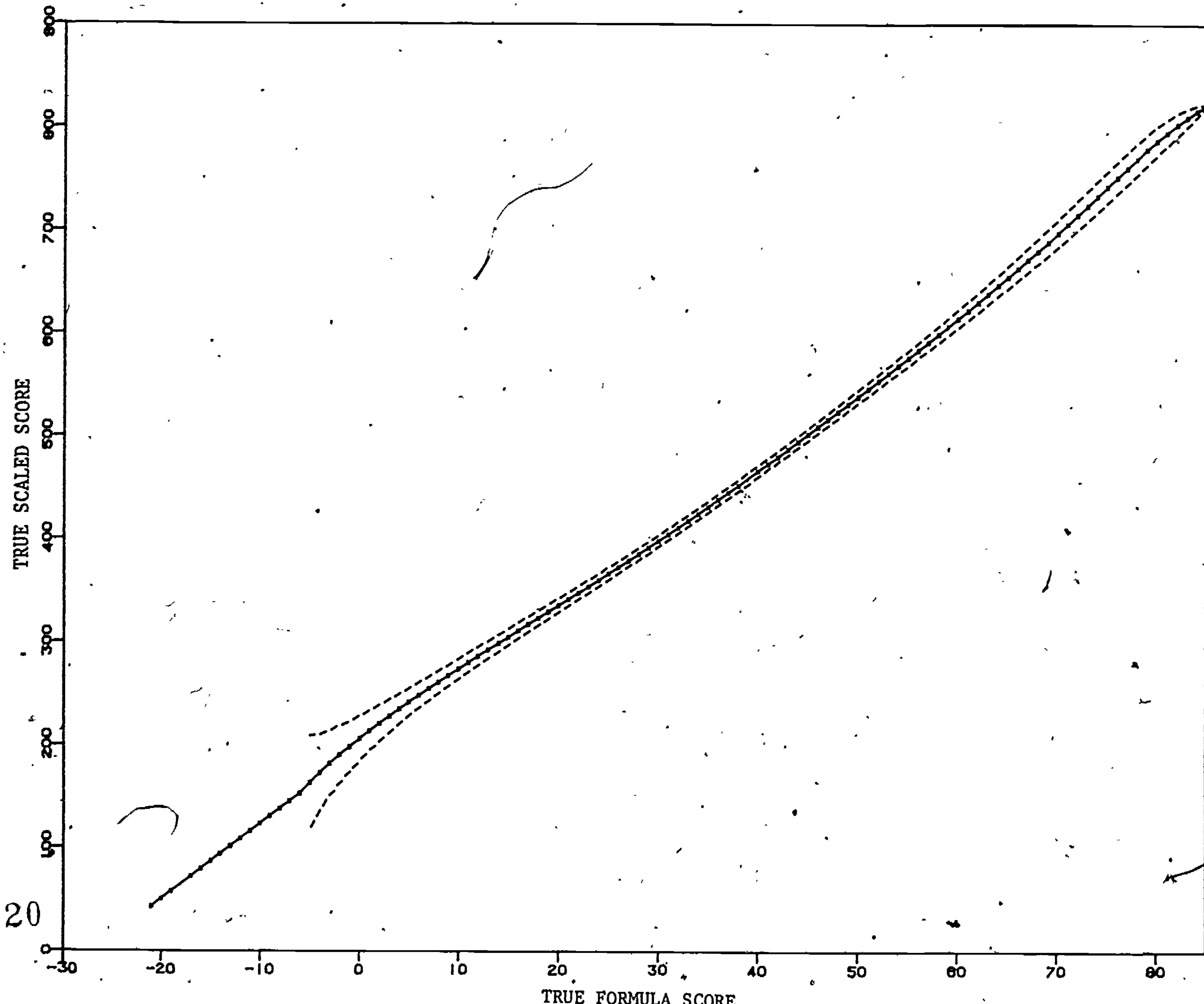


Figure 2. IRT equating of XSA2 formula score to VSA4 scaled score, with two-standard-error bounds.

Table 2

A Comparison of Equipercentile and IRT Equating and of Their Standard Scores

	Equipercentile Method		IRT Model	
XSA2 formula score	Equivalent scaled score	Standard error	Equivalent scaled score	Standard error
78.1	774	13.47	764	4.68
70.6	722	15.85	700	4.18
64.75	652	10.32	651	3.44
58.9	602	4.97	605	2.78
52.9	558	4.12	558	2.32
47.25	514	3.47	515	2.09
40.1	466	3.44	464	2.05
32.4	417	2.93	412	2.24
25.75	364	3.37	370	2.63
16.1	314	4.07	312	3.62
7.6	242	5.70	259	5.08
-3.75	195	7.85	175	12.49

Since SE_{XSA2} and SE_{VSA4} are estimated from unsmoothed data, the equipercentile standard errors in Table 2 fluctuate somewhat. Nevertheless, it is apparent that the equipercentile method has a much larger standard error above a scaled score of 450. For these data, the IRT method shows a larger standard error than the equipercentile method only when the formula score is negative.

The standard error of equipercentile equating could be reduced by smoothing the frequency distribution of raw scores before equating. Smoothing is undoubtedly desirable as a practical expedient; however, the choice of a smoothing formula is somewhat arbitrary and the smoothing is likely to prevent convergence of the estimated equating to its true value in large samples. Formulas for the standard errors of smoothed equipercentile equating are not presently available.

In order to determine the effect of using a shorter anchor test, every other item in the anchor test was discarded and the data reanalyzed on the basis of the remaining 20-item anchor test. The effect on the standard errors of IRT equating in shown in Table 3. The two equatings agree fairly well. At the point where the equating standard errors are a minimum, halving the length of the anchor test increases the standard error by a factor of about $\sqrt{2}$. At the other score points, the effect is less. Given standard errors like those in Table 2, it will now be possible to make a reasonable judgment as to the length necessary for an anchor test.

Table 3

IRT Equatings and Their Scaled-Score Standard Errors,

a Comparison of Results Using 20- and 40-Item Anchor Tests

XSA2 formula score	Length of Anchor Test: 20 Items		40 Items	
	Scaled score	Standard error	Scaled score	Standard error
80	787	5.9	780	4.5
70	698	5.3	695	4.1
60	615	3.9	613	2.9
50	540	3.0	536	2.2
40	467	2.7	463	2.0
30	399	3.0	397	2.4
20	336	3.9	335	3.2
10	274	5.4	275	4.6
0	206	9.9	206	8.4

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Chalfont, PA 18914

1 Dr. Robert Wisher
Code 309
Navy Personnel R & D Center
San Diego, CA 92152

1 Dr. Martin F. Wiskoff
Navy Personnel R & D Center
San Diego, CA 92152

1 Mr. Ted M. I. Yellen
Technical Information Office
Code 201
Navy Personnel R & D Center
San Diego, CA 92152

Army

1 Technical Director
U.S. Army Research Institute for the
Behavioral and Social Sciences
5001 Eisenhower Avenue
Alexandria, VA 22333

1 Dr. Myron Fischl
U.S. Army Research Institute for the
Social and Behavioral Sciences
5001 Eisenhower Avenue
Alexandria, VA 22333

1 Dr. Dexter Fletcher
U.S. Army Research Institute
5001 Eisenhower Avenue
Alexandria, VA 22333

1 COL Frank Hart
Army Research Institute for the
Behavioral & Social Sciences
5001 Eisenhower Blvd.
Alexandria, VA 22333

1 Dr. Michael Kaplan
U.S. Army Research Institute
5001 Eisenhower Avenue
Alexandria, VA 22333

1 Dr. Milton S. Katz
Training Technical Area
U.S. Army Research Institute
5001 Eisenhower Avenue
Alexandria, VA 22333

1 Dr. Harold F. O'Neil, Jr.
Attn: PERI-OK
Army Research Institute
5001 Eisenhower Avenue
Alexandria, VA 22333

1 Dr. Robert Sasmor
U.S. Army Research Institute for
the Social and Behavioral Sciences
5001 Eisenhower Avenue
Alexandria, VA 22333

1 Commandant
U.S. Army Institute of Administration
Attn: Dr. Sherrill
Ft. Benjamin Harrison, IN 46256

1 Dr. Frederick Steinheiser
Department of the Navy
Chief of Naval Operations
OP-113
Washington, DC 20350

1 Dr. Joseph Ward
U.S. Army Research Institute
5001 Eisenhower Avenue
Alexandria, VA 22333

Air Force

1 Air Force Human Resources Laboratory
AFHRL/MPD
Brooks Air Force Base, TX 78235

1 U.S. Air Force Office of
Scientific Research
Life Sciences Directorate
Bolling Air Force Base
Washington, DC 20332

1 Air University Library
AUL/LSE 76/443
Maxwell Air Force Base, AL 36112

1 Dr. Earl A. Alluisi
HQ, AFHRL (AFSC)
Brooks Air Force Base, TX 78235

1 Mr. Raymond E. Christal
AFHRL/MO
Brooks Air Force Base, TX 78235

1 Dr. Genevieve Haddad
Program Manager
Life Sciences Directorate
AFOSR
Bolling Air Force Base
Washington, DC 20332

1 Dr. Ross L. Morgan
AFHRL/LR
Wright-Patterson AFB, OH 45433

1 Research and Measurement Division
Research Branch, AFMPC/MPCYPR
Randolph Air Force Base, TX 78148

1 Dr. Malcolm Ree
AFHRL/MP
Brooks Air Force Base, TX 78235

1 Dr. Marty Rockway
Technical Director
AFHRL(OT)
Williams Air Force Base, AZ 58224

Marines

1 Dr. H. William Greenup
Education Advisor (E031)
Education Center, MCDEC
Quantico, VA 22134

1 Director, Office of Manpower Utilization
HQ, Marine Corps (MPU)
BCB, Building 2009
Quantico, VA 22134

1 MAJ Michael L. Patrow, USMC
Headquarters, Marine Corps
Code MPI-20
Washington, DC 20380

1 Dr. A. L. Slafkosky
Scientific Advisor
Code RD-1
HQ, U.S. Marine Corps
Washington, DC 20380

Coast Guard

1 Mr. Thomas A. Warm
U.S. Coast Guard Institute
P.O. Substation 18
Oklahoma City, OK 73169

Other DoD

1 DARPA
1400 Wilson Boulevard
Arlington, VA 22209

12 Defense Technical Information Center
Cameron Station, Building 5
Attn: TC
Alexandria, VA 22314

1 Dr. William Graham
Testing Directorate
MEPCOM/MEPCT-P
Ft. Sheridan, IL 60037

1 Military Assistant for Training
and Personnel Technology
Office of the Under Secretary of
Defense for Research and Engineering
Room 3D129, The Pentagon
Washington, DC 20301

1 Dr. Wayne Sellman
Office of the Assistant Secretary
of Defense (MRAL)
2B269 The Pentagon
Washington, DC 20301

Civil Government

1 Dr. Susan Chipman
Learning and Development
National Institute of Education
1200 19th Street, NW
Washington, DC 20208

1 Mr. Richard McKillip
Personnel R & D Center
Office of Personnel Management
1900 E Street, NW
Washington, DC 20415

1 Dr. Arthur Melmed
National Institute of Education
1200 19th Street, N.W.
Washington, DC 20208

1 Dr. Andrew R. Molnar
Science Education Development
and Research
National Science Foundation
Washington, DC 20550

1 Dr. Vern W. Urry
Personnel R & D Center
Office of Personnel Management
1900 E Street, NW
Washington, DC 20415

1 Dr. Joseph L. Young, Director
Memory and Cognitive Processes
National Science Foundation
Washington, DC 20550

Non-Government

1 Dr. Erling B. Andersen
Department of Statistics
Studiestraede 6
1455 Copenhagen
DENMARK

1 Psychological Research Unit
Department of Defense (Army Office)
Campbell Park Offices
Canberra, ACT 2600
AUSTRALIA

1 Dr. Alan Baddeley
Medical reserch Council
Applied Psychology Unit
15 Chaucer Road
Cambridge CB2 2EF
ENGLAND

1 Dr. Isaac Bejar
Educational Testing Service
Princeton, NJ 08541

1 Dr. Menucha Birenbaum
School of Education
Tel Aviv University
Tel Aviv, Ramat Aviv 69978
ISRAEL

1 Dr. Werner Birke
DezWPs im Streitkraefteamt
Postfach 20 50 3
D-5300 Bonn 2
WEST GERMANY

1 Dr. R. Darrell Bock
Department of Education
University of Chicago
Chicago, IL 60637

1 Liaison Scientists
Office of Naval Research
Branch Office, London
Box 39
FPO, NY 09510

1 Dr. Robert Brennan
American College Testing Programs
P.O. Box 168
Iowa City, IA 52240

1 Dr. John B. Carroll
Psychometric Laboratory
University of North Carolina
Davie Hall 013A
Chapel Hill, NC 27514

1 Charles Myers Library
Livingstone House
Livingstone Road
Stratford
London E15 2LJ
ENGLAND

1 Dr. Kenneth E. Clark
College of Arts and Sciences
University of Rochester
River Campus Station
Rochester, NY 14627

1 Dr. Norman Cliff
Department of Psychology
University of Southern California
University Park
Los Angeles, CA 90007

1 Dr. William E. Coffman
Director, Iowa Testing Programs
334 Lindquist Center
University of Iowa
Iowa City, IA 52242

1 Dr. Allan M. Collins
Bolt, Beranek, & Newman, Inc.
50 Moulton Street
Cambridge, MA 02138

1 Dr. Meredith P. Crawford
American Psychological Association
1200 17th Street, N
Washington, DC 20036

1 Dr. Hans Crombag
Education Research Center
University of Leyden
Boerhaavelaan 2
2334 EN Leyden
THE NETHERLANDS

1 Dr. Fritz Drasgow
Yale School of Organization and Management
Yale University
Box 1A
New Haven, CT 06520

1 LCOL J. C. Eggenberger
Directorate of Personnel Applied Research
National Defence Hq.
101 Colonel By Drive
Ottawa, KIA OK2
CANADA

1 Dr. Benjamin A. Fairbank, Jr.
McFann-Gray and Associates, Inc.
5825 Callaghan
Suite 225
San Antonio, TX 78228

1 Dr. Leonard Feldt
Lindquist Center for Measurement
University of Iowa
Iowa City, IA 52242

1 Dr. Richard L. Ferguson
The American College Testing Program
P.O. Box 168
Iowa City, IA 52240

1 Dr. Victor Fields
Department of Psychology
Montgomery College
Rockville, MD 20850

1 Univ. Prof. Dr. Gerhard Fischer
Psychologisches Institut der
Universitat Wien
Liebiggasse 5/3
A 1010 Wien
AUSTRIA

1 Prof. Donald Fitzgerald
University of New England
Armidale, New South Wales 2351
AUSTRALIA

1 Dr. Edwin A. Fleishman
Advanced Research Resources Organization
Suite 900
4330 East West Highway
Washington, DC 20014

1 Dr. John R. Frederiksen
Bolt, Beranek, and Newman
50 Moulton Street
Cambridge, MA 02138

1 Dr. Robert Glaser
LRDC
University of Pittsburgh
3939 O'Hara Street
Pittsburgh, PA 15213

1 Dr. Bert Green
Department of Psychology
Johns Hopkins University
Charles and 34th Streets
Baltimore, MD 21218

1 Dr. Ron Hambleton
School of Education
University of Massachusetts
Amherst, MA 01002

1 Dr. Lloyd Humphreys
Department of Psychology
University of Illinois
Champaign, IL 61820

1 Library
HumRRO/Western Division
27857 Berwick Drive
Carmel, CA 93921

1 Dr. Steven Hunka
Department of Education
University of Alberta
Edmonton, Alberta
CANADA

1 Dr. Earl Hunt
Department of Psychology
University of Washington
Seattle, WA 98105

1 Dr. Jack Hunter
2122 Coolidge Street
Lansing, MI 48906

1 Dr. Huynh Huynh
College of Education
University of South Carolina
Columbia, SC 29208

1 Mr. Marlin Kroger
1117 Via Goleta
Palos Verdes Estates, CA 90274

1 Dr. Michael Levine
Department of Educational Psychology
210 Education Building
University of Illinois
Champaign, IL 61801

1 Dr. Charles Lewis
Faculteit Sociale Wetenschappen
Rijksuniversiteit Groningen
Oude Boteringestraat 23
9712GC Groningen
NETHERLANDS

1 Dr. Robert Linn
College of Education
University of Illinois
Urbana, IL 61801

1 Dr. James Lumsden
Department of Psychology
University of Western Australia
Nedlands, Western Australia 6009
AUSTRALIA

1 Dr. Gary Marco
Educational Testing Service
Princeton, NJ 08541

1 Dr. Scott Maxwell
Department of Psychology
University of Houston
Houston, TX 77004

1 Dr. Samuel T. Mayo
Loyola University of Chicago
820 North Michigan Avenue
Chicago, IL 60611

1 Dr. Allen Munro
Behavioral Technology Laboratories
1845 Elena Avenue
Fourth Floor
Redondo Beach, CA 90277

1 Dr. Melvin R. Novick
356 Lindquist Center for Measurement
University of Iowa
Iowa City, IA 52242

1 Dr. Jesse Orlansky
Institute for Defense Analyses
400 Army Navy Drive
Arlington, VA 22202

1 Dr. Wayne M. Patience
American Council on Education
GED Testing Service, Suite 20
One Dupont Circle, NW
Washington, DC 20036

1 Dr. James A. Paulson
Portland State University
P.O. Box 751
Portland, OR 97207

1 Mr. Luigi Petrullo
2431 North Edgewood Street
Arlington, VA 22207

1 Dr. Diane M. Ramsey-Klee
R-K Research and System Design
3947 Ridgemont Drive
Malibu, CA 90265

1 Mr. Minrat M. L. Rauch
P II 4
Bundesministerium der Verteidigung
Postfach 1328
D-53 Bonn 1
GERMANY

1 Dr. Mark D. Reckase
Educational Psychology Department
University of Missouri-Columbia
4 Hill Hall
Columbia, MO 65211

1 Dr. Leonard L. Rosenbaum, Chairman
Department of Psychology
Montgomery College
Rockville, MD 20850

1 Dr. Ernst Z. Rothkopf
Bell Laboratories
600 Mountain Avenue
Murray Hill, NJ 07974

1 Dr. Lawrence Rudner
403 Elm Avenue
Takoma Park, MD 20012

1 Dr. J. Ryan
Department of Education
University of South Carolina
Columbia, SC 29208

1 Prof. Fumiko Samejima
Department of Psychology
University of Tennessee
Knoxville, TN 37916

1 Dr. Frank L. Schmidt
Ddepartment of Psychology
Building GG
George Washington University
Washington, DC 20052

1 Dr. Robert J. Seidel
Instructional Technology Group
HumRRO
300 North Washington Street
Alexandria, VA 22314

1 Committee on Cognitive Research
c/o Dr. Lonnie R. Sherrod
Social Science Research Council
605 Third Avenue
New York, NY 10016

1 Dr. Kazuo Shigemasu
University of Tohoku
Department of Educational Psychology
Kawauchi, Sendai 980
JAPAN

1 Dr. Edwin Shirkey
Department of Psychology
University of Central Florida
Orlando, FL 32816

1 Dr. Richard Snow
School of Education
Stanford University
Stanford, CA 94305

1 Dr. Robert Sternberg
Department of Psychology
Yale University
Box 11A, Yale Station
New Haven, CT 06520

1 Dr. Albert Stevens
Bolt, Beranek, and Newman, Inc.
50 Moulton Street
Cambridge, MA 02138

1 Dr. Hariharan Swaminathan
Laboratory of Psychometric and
Evaluation Research
School of Education
University of Massacuusetts
Amherst, MA 01003

1 Dr. Kikumi Tatsuoka
Computer Based Education Research
Laboratory
252 Engineering Research Laboratory
University of Illinois
Urbana, IL 61801

1 Dr. David Thissen
Department of Psychology
University of Kansas
Lawrence, KS 66044

1 Dr. Robert Tsutakawa
Department of Statistics
University of Missouri
Columbia, MO 65201

1 Dr. David Vale
Assessment Systems
Corporation
2395 University Avenue
Suite 306
St. Paul, MN 55114

1 Dr. Howard Wainer
Educational Testing Service
Princeton, NJ 08541

1 Dr. Thomas Wallsten
Psychometric Laboratory
Davie Hall 013A
University of North Carolina
Chapel Hill, NC 27514

1 Dr. Phyllis Weaver
Graduate School of Education
Harvard University
200 Larsen Hall, Appian Way
Cambridge, MA 02138

1 Dr. David J. Weiss
N660 Elliott Hall
University of Minnesota
75 East River Road
Minneapolis, MN 55455

1 Dr. Susan E. Whitely
Psychology Department
University of Kansas
Lawrence, KS 66044

1 Dr. Wolfgang Wildgrube
Streitkraefteamt
Box 20 50 03
D-5300 Bonn 2
WEST GERMANY